EXHIBIT 22
[FILED UNDER SEAL]

FAMILY	1 if sample respondent reports that "raise a family" is an essential goal and value; 0 otherwise.
DRIVE	1 if rates self in the highest 10% in terms of "drive to achieve"; 0 otherwise.
SUCCEED	1 if "be successful in own business" is an essential goal and value; 0 otherwise.
WELLOFF	1 if "be well off financially' is an essential goal and value; 0 otherwise.
INCOME	Current (1980) annual income before taxes: 1 = no income; 2 = \$1 to \$6,999; 3 = \$7,000 to \$9,999; 4 = \$10,000 to \$14,999; 5 = \$15,000 to \$19,999; 6 = \$20,000 to \$24,999; 7 = \$25,000 to \$29,999; 8 = \$30,000 to \$34,999; 9 = \$35,000 to \$39,999; 10 = \$40,000 or more.

Note: Descriptive statistics are available from the authors.

PRICE FIXING: THE PROBABILITY OF GETTING CAUGHT

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Abstract—We estimate the probability that a price fixing conspiracy will be indicted by federal authorities to be at most between 0.13 and 0.17 in a given year. Our estimate is based on conspiracy durations calculated from data reported for a large sample of DOJ cases, and a statistical birth and death process model describing the onset and duration of conspiracies.

I. Introduction

The economic theory of collusion (e.g., see Stigler (1964)) suggests that a firm's decision to participate in a price fixing conspiracy is based substantially on a rational calculation of associated benefits and costs. One cost is the possible penalties arising from getting caught. The greater the probability of getting caught, the greater these penalties loom in the prospective conspirators' calculations, and, ceteris paribus, the less likely they are to collude. The probability of getting caught therefore is a measure of the deterrent effect of antitrust enforcement, and should be inversely related to the number of price fixing conspiracies attempted.

Our paper is the first to estimate this probability. We also estimate the number of conspiracies (eventually caught) active at one time. The results are based on approximate conspiracy durations calculated from data reported for a large sample of U.S. Department of Justice price fixing indictments. The intuition behind our approach is simple. If the distribution of conspiracy durations is (e.g.) characterized by many short-lived conspiracies and few long-lived ones, then the probability of getting caught must be high, and, given the catch-rate, the total number of active conspiracies (eventually caught) must be low. We propose a statistical model to describe the onset and duration of such conspiracies, a simple birth-and-death process model, and use maximum likelihood methods to estimate model parameters. These parameter estimates in turn are the basis for our estimates of the number of active conspiracies (eventually caught) and the probability of getting caught.

The total population of active conspiracies can be partitioned into two subpopulations, one containing conspiracies which are eventually caught and another containing those which are not. Our sample is from the first subpopulation, and therefore our statistical inferences relate directly to that population only. No data exist regarding the second population. Nevertheless, one can extrapolate to the second some information inferred regarding the first.

II. The Model

Suppose at time $t = 0$ there exist no price conspiracies. At any time $t > 0$, let there be $N(t)$ active ("alive") conspiracies. The changes in $N(t)$ can be described by a birth-and-death process, a continuous time Markov chain, as follows. In a short interval of time from t to

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$t + h$, $N(t)$:

(1) changes from $N(t)$ to $N(t) + 1$ with probability $\theta h + o(h)$;[1]

(2) changes from $N(t)$ to $N(t) - 1$ with probability $N(t) \cdot \lambda h + o(h)$; or

(3) remains unchanged with probability $1 - \theta h - \lambda h + o(h)$.

No other changes are possible, for h sufficiently small.

In such a process, the lifetimes L_t and interarrival times A_t between the births of successive conspiracies are independently and exponentially distributed with means λ^{-1} and θ^{-1}, respectively; and $N(T)$, the number alive at a particular time T, has a Poisson distribution with mean $\theta_T = (\theta/\lambda)\{1 - e^{-\lambda T}\}$.[2] As T becomes large, the effect of the start-up time decreases and the process approaches a steady state in which the expected number of conspiracies alive is θ/λ. If the model is approximately correct, our initial questions may be answered by estimating model parameters θ and λ. Section V and the appendix contain the details of that estimation.

The birth-and-death process was originally suggested by the rough exponentiality of lifetime distributions (see table 1) and because its lack of memory property makes it mathematically tractable. A plausible alternative model may be derived by assuming that (in the language of reliability theory) the hazard rate, the probability of a conspiracy's being caught in the interval t to $t + h$, should be roughly proportional to the evidence $E(t)$ available at time t, given that it has not been caught earlier. The resulting lifetime distribution is[3]

$$P\{L \le u\} = 1 - \exp\left\{-\int_0^u E(t)\,dt\right\}.$$

When $E(t)$ is a constant, this formula gives the exponential distribution, and the probability of getting caught in the next interval (given conspiracy existence at the end of the current interval) is independent of conspiracy duration, though the *cumulative* probability of having been caught does increase with age.

When $E(t)$ is proportional to $t^{\alpha-1}$ for some parameter α, the resulting distribution is the Weibull distribution.[4] The exponential distribution is a special case of the Weibull, when $\alpha = 1$ (constant hazard rate). It seems plausible that the hazard rate in our case would be constant or increasing ($\alpha > 1$). In section IV we show that for our data the exponential fit is reasonably good; the best fitting Weibull has $\alpha = 1.1$, very close to the exponential. Deviations from a constant hazard rate appear to be very small.[5]

Regarding other model assumptions, statistical independence seems reasonable. Since the birth and continued existence of conspiracies are presumably private matters, there is little reason to think that the onset or duration of one conspiracy should be influenced by another. The assumption of exponential inter-arrival times is a bit harder to justify. This implies that the "propensity" of society to spawn a conspiracy is approximately constant over time. While a priori arguments on this point are unclear, our sample conspiracies appear consistent with this assumption. Finally, our model assumes a steady state, i.e., the parameters θ and λ do not vary over time. This assumption is questionable, since presumably they *could* vary over time. On the other hand, the model derived with constant parameters appears to fit the data.

III. The Data

Our sample of horizontal price fixing conspiracies is selected from the Commerce Clearing House Trade Regulation Reports (the "Bluebook") for the period 1961 to 1988. The Bluebook presents summary histories of DOJ price fixing cases, including specific allegations and outcomes. The text material and index are arranged by conspiracy, rather than by case, with often two cases (e.g., one civil and one criminal) and sometimes more for the same conspiracy. All conspiracies listed in the Bluebook index for the sample period under "price fixing" were examined, except those involving retail price maintenance and monopsony. Dismissed cases or those ending in acquittals were dropped, as were cases with unreported outcomes.

We treat *nolo contendere* ("no contest") pleas as equivalent to guilty pleas and convictions in constituting evidence of a conspiracy.[6] This is an important assumption since the great majority of cases are settled by such pleas. Some argue that innocent firms frequently plead *nolo contendere* to avoid costly litigation.[7]

[1] The notation $o(h)$ means something which when divided by h tends to zero as h tends to 0. In this case, it means the probability is approximately θh, and the approximation is valid for small values of the interval width h.

[2] The definitions and results here for Markov chains and for the birth-and-death process in particular are taken from Feller (1957, chapter XVII, pp. 413–414 and p. 435); and Karlin (1966, chapter 7, pp. 196–197).

[3] E.g., see Karlin (1966), exercise 12, p. 211.

[4] E.g., see Mann et al. (1974), chapters 4–6.

[5] This is consistent with the suggestions by some (e.g., Posner (1970)) that the natural breakdown of conspiracies often leads to indictments. For example, if breakdowns are caused by exogenous shocks (e.g., demand shifts), the probability of breakdown in period $n + 1$ would be independent of the number of prior periods n of successful conspiracy.

[6] Hay and Kelley (1974) make this assumption in selecting their sample.

[7] We can't know how many innocent firms plead *nolo contendere*, if any. The presence of such firms could affect our analysis in two ways. First, if the distribution of alleged conspiracy durations for such firms differed from that of colluders, our observed sample distribution containing both types would not reflect the actual colluders' distribution. Second, our sample size would be exaggerated causing an upward bias in common measures of statistical significance.

However, price fixing cases are relatively straightforward and cheap, at least as compared to other types of antitrust cases which involve complex legal and economic theories and require extensive and detailed analysis of firm operations and market conditions. A more likely explanation is that *nolo contendere* pleas allow firms to avoid an official court determination of guilt. This substantially reduces the likelihood of successful follow-on private treble damage suits which must therefore prove guilt, in addition to damages (e.g., see Werden (1988)).

For our purposes, the critical Bluebook information concerns conspiracy duration. All summaries include the date of indictment, beyond which the conspiracy presumably no longer exists. Most summaries contain a more or less specific date described as the conspiracy beginning. Cases with unreported conspiracy initiation dates are excluded from our sample. In addition, many cases report a third date, prior to the indictment date, at which the conspiracy may have ended. An approximate duration can therefore be calculated as the interval between the "begin" date and the indictment or "end" date.

The indictment date is precise (month, day, and year are reported) and unambiguous; however, the other dates generally are not. With few exceptions, the most precise identification is to the month, e.g., a statement that the "conspiracy began in April 1960" (case no. 1592). In other cases, only a year is given, e.g., the "price fixing conspiracy began in 1957 and continued until sometime in 1961" (no. 1643). In still others, periods greater than a month but less than a year are given, e.g., the conspiracy began "in the fall of 1966" (no. 1953). A second source of ambiguity regarding the reported non-indictment dates is that in many cases the implied begin (end) date may be only an upper (lower) bound on the true date. In several cases, the words "at least" accompany the reported period. For example, case no. 2179 reports that the defendants conspired "since at least January 1966." In some cases, it is explicit that the reported period is a bound, e.g., the defendants are charged with conspiring "since before 1963" (no. 2368). A third problem is that the begin and end dates identified in the Bluebook summaries may simply represent the earliest and latest points at which evidence of the conspiracy is available. If this is true, our conspiracy durations are all strict lower bounds.

We use two methods to calculate conspiracy durations. Each accepts the begin date as valid, and for the foregoing reasons should be interpreted as a lower bound. The first method defines duration to be the period from the earliest begin day to the indictment day (*DUR*1).[8] For example, if the Bluebook indicates that the conspiracy began in 1971 and continued to at least October 1975, the earliest begin day is January 1, 1971. *DUR*1 is the longest duration which can be calculated with available data. The second method defines the duration to be the period from the latest begin day to the earliest end day (*DUR*2). The indictment day is used if no end day is given. In the above example, the latest begin day is December 31, 1971, and the earliest end day is October 1, 1975. *DUR*2 is the shortest duration which can be calculated with available data. These two measures cover the range of duration values which can be generated by alternative calculation procedures, and each is defined to include all cases in the sample.

TABLE 1.—FREQUENCY DISTRIBUTIONS AND SUMMARY STATISTICS FOR CONSPIRACY DURATION VARIABLES

	Frequencies	
Interval (Yrs.)	*DUR*1[a]	*DUR*2[b]
0–2 Years	12	55
2–4	42	42
4–6	42	31
6–8	25	15
8–10	21	8
10–12	12	14
12–14	12	10
14–16	5	3
16–18	9	2
18–20	0	2
> 20	4	2
Sample Size	184	184
Mean	7.27 yrs.	5.23 yrs.
Median	5.80	3.61
Minimum	0.33	0.01
Maximum	40.8	39.8

[a] Calculated as interval between earliest begin date and indictment date.
[b] Calculated as interval between latest begin date and earliest end date, if reported, or indictment date, if not.

Summary statistics and frequency distributions for the duration measures are reported in table 1. Our sample contains 184 conspiracies. *DUR*1 has a mean and median of 7.27 years and 5.80 years, respectively, and varies from 0.33 years to 40.8 years. *DUR*2 has a mean and median of 5.23 and 3.61, respectively, and varies from 0.01 to 39.8.[9]

The Bluebook also reports two variables which economic theory suggests should be related to conspiracy duration. These are the number of firms involved in the conspiracy (*NF*, reported for 101 conspiracies) and their market share (*SHR*, reported for 27 conspiracies).[10] Conspiracy duration should be longer

[8] The time function in our statistical software package requires that dates be specified to the day.

[9] Posner (1970) reports a mean conspiracy duration of 6 years for his sample of DOJ indictments covering 1960–69, based also on Bluebook summaries. He does not describe how he computes durations.

[10] Information on the number of involved firms can be missing because (1) in many cases the Bluebook summary indicates the presence of unindicted and unnamed co-conspirators, and (2) some cases involve trade associations whose member firms are not individually indicted or identified.

FIGURE 1.—EXPONENTIAL FIT OF CONSPIRACY DURATIONS

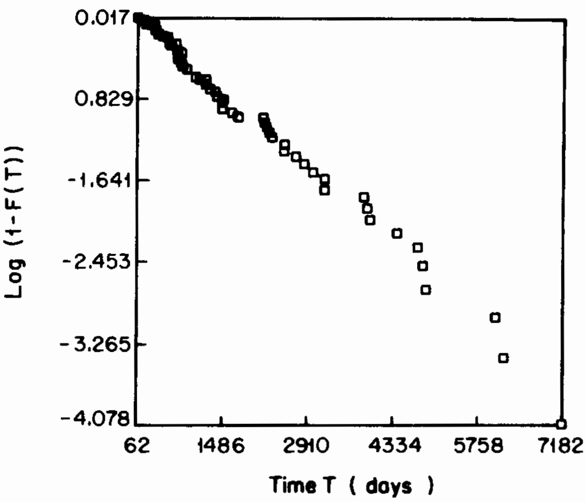
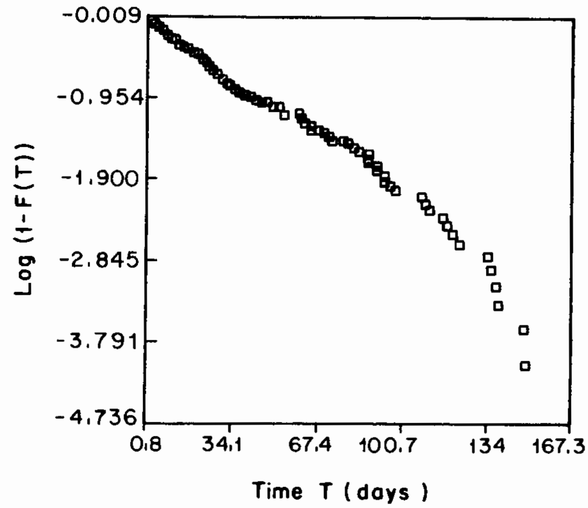


FIGURE 2.—EXPONENTIAL FIT OF CONSPIRACY INTERARRIVAL TIMES



the fewer the involved firms and the larger their market share, ceteris paribus. While these hypotheses are clear for conspiracies terminating for "natural" causes (e.g., irreconcilable differences among conspirators), it is less clear that they apply to conspiracies terminating because of prosecution. Be that as it may, neither variable is correlated with our conspiracy duration measures.[11] It may be that conspiracies with unfavorable values of *NF* or *SHR* are launched *only* if there are other offsetting favorable factors about which we have no information. We therefore feel justified in treating conspiracy duration as a random variable.

IV. Model Fit

To see how close our assumption of exponentiality fits the interarrival time and duration data, we look at the empirical cumulative distribution function, $F(x)$ = (number of observations $\leq x$)/(total number of observations). For an exponential distribution, $\log(1 - F)$ should be approximately linear in x. In addition, we may test the null hypothesis that the distribution is exponential versus the alternative that it is Weibull with a non-unity shape parameter using a likelihood ratio test.

For conspiracy durations, data are used from 59 cases with reasonably exact dates. The remaining cases are dropped to avoid ambiguities arising from the use of lower bounds. The start and end dates are first taken as the midpoint of the interval for each specified by the data, except that the indictment date is used if no other end date is specified. The durations are then computed from the start and end dates. The resulting plot of $\log(1 - F)$ versus x (conspiracy duration) is given in figure 1, and is approximately linear. An OLS regression of $\log(1 - F)$ on duration yields an $R^2 = 0.994$. When fitting a Weibull distribution by maximum likelihood,[12] the estimated shape parameter is 1.092 (exponential is 1.00). The likelihood ratio test of exponentiality gives a chi-square statistic of 0.74 on 1 degree of freedom, which would not justify rejecting the hypothesis of exponentiality at reasonable significance levels. The duration data seem reasonably consistent with the exponential assumption.

For interarrival times, the data for 114 conspiracies starting between 1962 and 1976 are used.[13] The start dates are taken randomly within the interval specified, and the resulting interarrival times are computed.[14] Figure 2 plots $\log(1 - F)$ versus x (interarrival time), and shows an approximate linear trend ($R^2 = 0.985$). Exponentiality would not be rejected statistically in favor of a Weibull distribution in this case either. The Weibull shape parameter is 0.984, and the chi-squared statistic is 0.05 on 1 degree of freedom. We adopt the exponential model, tentatively, as it appears to describe the bulk of the data satisfactorily and it is mathematically tractable.

[11] Posner (1970) reports similar results for the number of involved firms and duration in his sample covering 1950–69. The presence of innocent *nolo contendere* firms might bias the correlations toward zero.

[12] E.g., see Mann et al. (1974).

[13] It's necessary to choose an interval inside the range of indictments considered, to minimize situations in which conspiracies and their starting times are missed because their indictment date fell outside the 1961–1988 range.

[14] Random assignment minimizes the occurrence of zero interarrival times caused by the midpoint assignment convention. For example, three conspiracies alleged to begin in the same month would, by the midpoint convention, have identical start dates and thus zero interarrival times.

FIGURE 3.—NUMBER OF OBSERVED CONSPIRACIES ALIVE AT VARIOUS TIMES

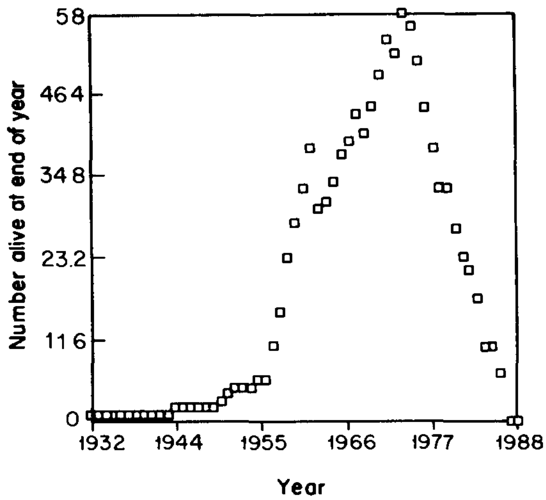


TABLE 2.—PARAMETER ESTIMATES

Duration Measure	"Deaths" $\lambda^a(1/\lambda)^b$	"Births" $\theta^c(1/\theta)^d$	Number "Alive" θ/λ
*DUR*1	.000377 (2655)	.0187 (53.6)	49.5
*DUR*2	.000524 (1910)	.0187 (53.6)	35.6

[a] Probability of conspiracy "death" on a given day
[b] Mean conspiracy duration in days.
[c] Probability of conspiracy "birth" on a given day
[d] Mean interarrival time in days

V. Parameter Estimates

The parameters θ and λ are estimated by the maximum likelihood method, as derived in the appendix. The method requires us to specify not only the sample period (January 1961 to December 1987, or 9,860 days), but also how long the process had continued before we began observing (or else assume that this time interval T is "large"). We don't know how long there have been price conspiracies, but they have likely existed for many years prior to 1961. Calculations using T values ranging from 25 to 100 years reveal that the differences arising from assuming that T is "large" are negligible. Therefore all further reported results assume that T is "large," i.e., that the process in 1961 had reached a steady state.

The resulting estimates of θ and λ are shown in table 2 for our two duration measures.[15] The average interarrival time $(1/\theta)$ is about 50 days, or $\theta = .02$. This is the same for both *DUR*1 and *DUR*2 since the time between conspiracy births is unrelated to conspiracy duration. With *DUR*1, the probability of getting caught (λ) is about 3.8×10^{-4} per day (0.128 per year). With *DUR*2, the probability of getting caught is about 5.2×10^{-4} per day (0.174 per year). As noted earlier, *DUR*1 and *DUR*2 are best thought of as lower bounds, implying that our λ estimates are upper bounds. The average number of conspiracies alive at one time (θ/λ) is 49.5 for *DUR*1, and 35.6 for *DUR*2. Since λ is an upper bound, θ/λ is a lower bound.

The precision of these estimates can be gauged roughly by using log-likelihood values. If we take as rough intervals for θ and λ those values (based on *DUR*1) which give a log-likelihood within about 2 of its maximum, we may take θ as between approximately 0.016 and 0.022, and λ between approximately 0.00032 and 0.00045. This implies a range of about 36 to 69 for the average number of conspiracies "alive" at time θ/λ, with a value of about 50 the most likely. The above specified range for λ implies a probability of getting caught in any given year of between 0.110 and 0.151, with 0.128 being most likely.

The model predicts that the number of conspiracies alive at any time T should fluctuate around an expected value of θ/λ. Because we observe the process only from times T_1 to T_2 $(T_1 < T < T_2)$, and not forever, it will not appear that quite that many are alive at T, since those alive at both T_1 and T_2 won't be observed. The expected number that we will *observe* to be alive at time T (see appendix), given that we observe only until T_2, is $(\theta/\lambda)(1 - \exp\{T_2 - T\})$. The closer we get to time T_2, the more we expect the number observed alive to be attenuated by the exclusion of those not yet caught. Figure 3 plots the number actually observed alive (based on *DUR*1) at various times, and is consistent with expectations.

VI. Conclusions

According to our model, the average number of (eventually caught) conspiracies alive at any given time is at least 36 to 50. The average conspiracy lasts about five to seven years, and a new one is born about every 54 days (about seven per year). The probability of getting caught in a given year is at most between 0.13 and 0.17. The data's close fit to our exponential model suggests that the probability of a conspiracy getting caught in a given period, given that it is alive at the beginning of the period, is not related to its prior duration.

To what extent might these results apply to *uncaught* conspiracies? While such conspiracies presumably exist, there appears to be no way of estimating their number. However, if the life of a caught conspiracy is typically *no longer* than that of an uncaught conspir-

[15] Results for several other measures based on using upper and lower bounds range from $1/\lambda = 2187$ to as high as $1/\lambda = 3471$ days. The range given here (1902 to 2654) seems reasonable.

acy, then our duration estimates apply (also as lower bounds), and our probability estimate is an upper bound on the probability of uncaught conspiracy failure (i.e., demise by natural causes).

The probability of getting caught provides one piece of the expected cost calculations of prospective colluders (see Introduction). Other pieces include possible fines and treble-damage awards, and the likelihood of successful follow-on damage suits.[16] Putting these pieces together is a task for future research.

APPENDIX

The estimation methods and formulas used in the text combine standard features of birth-and-death processes (based on exponential assumptions) with ideas on censored data from reliability theory. The mathematical techniques for birth and death processes are illustrated in Karlin (1966) and Feller (1957), for example, and those for reliability theory in Mann et al. (1974). As the particular combination of techniques might not be familiar to readers of this *Review*, the derivations are sketched briefly here. Further details are available from the authors on request.

We assume the birth-and-death process begins at time $T = 0$, and we observe it from time T_1 to T_2 $(0 < T_1 < T_2)$. We *observe* only those conspiracies which "die" in the interval $[T_1, T_2]$, though others (unobserved) may have been born and died before T_1 or may still be alive at time T_2, and will subsequently be caught.

Using the independent exponential natures of the distributions of interarrival times and durations, it then follows that the probability of observing n conspiracies of durations $L_1, \dots, L_n$ is

$$\theta^n \lambda^n \exp[-\theta(T_2 - T_1)] \exp\left[-\lambda \sum L_i\right] \exp[w] \qquad \text{(A.1)}$$

where

$$w = (\theta/\lambda)\exp[-\lambda T_1]\{1 - \exp[-\lambda(T_2 - T_1)]\}$$

and θ^{-1} and λ^{-1} are the mean interarrival time and duration, respectively, of the conspiracies.

Equation (A.1) is the likelihood to be maximized as a function of θ and λ. If T_1 is large (the process has been going on a long time), then e^w is about 1, and (A.1) is maximized by $\theta^{-1} = (T_2 - T_1)/n$ and $\lambda^{-1} = (\Sigma L_1)/n$. If T_1 is not assumed large, a numerical technique may be used to maximize (A.1) directly. The Newton-Raphson technique is appropriate, as the second derivatives of (A.1) are negative. The optimal values of θ and λ are related by

$$\theta^{-1} = (T_2 - T_1)/n - (n\lambda)^{-1}\exp[-\lambda T_1] \times\{1 - \exp[-\lambda(T_2 - T_1)]\}$$

which shows how θ is inflated when T_1 is not negligible.

When the observed lifetimes L_i are only approximate, (A.1) can be modified as follows. Replace each factor $\lambda \exp[-\lambda L_i]$ by

$\{1 - \exp[-\lambda L_i]\}$ if L_i is an upper bound to the true lifetime;
$\exp[-\lambda L_i]$ if L_i is a lower bound to the true lifetime; or
$\exp[-\lambda L_i] - \exp[-\lambda U_i]$ if the true lifetime is bounded between two values L_i and U_i.

The thus-modified (A.1) may also be maximized via Newton-Raphson techniques.

Finally, note that since we observe the process only from T_1 to T_2, we expect the number of conspiracies *observed* alive at time t $(t > T_1)$ to taper off as $t \to T_2$, accounting for those processes which *were* alive at time t, but were still alive at time T_2, and hence haven't been discovered yet. This may be shown to follow a Poisson process with mean

$$E_t = (\theta/\lambda)\{1 - \exp[-\lambda t]\}\{1 - \exp[-\lambda(T_2 - t)]\}.$$

If T_1 is large, $E_t \sim (\theta/\lambda)\{1 - \exp[-\lambda(T_2 - t)]\}$, showing the expected behavior as illustrated in figure 3.

[16] For evidence on these issues, see Salop and White (1988), Werden (1988), Bosch and Eckard (1991), and references cited therein. Bosch and Eckard provide evidence that stock markets consider price fixing to be profitable net of expected legal costs.

REFERENCES

Bosch, Jean-Claude, and E. Woodrow Eckard, "The Profitability of Price Fixing: Evidence from Stock Market Reaction to Federal Indictments," this REVIEW 73 (May 1991), 309–317.

Feller, William, *An Introduction to Probability Theory and Its Applications*, Volume I, Second Edition (New York: John Wiley and Sons, Inc. 1957).

Hay, George A., and Daniel Kelley, "An Empirical Survey of Price Fixing Conspiracies," *Journal of Law and Economics* (Apr. 1974), 13–38.

Karlin, Samuel, *A First Course in Stochastic Processes* (New York: Academic Press, 1966).

Mann, N. R., R. E. Schafer, and N. D. Singpurwalla, *Methods for Statistical Analysis of Reliability and Life Data* (New York: John Wiley and Sons, 1974).

Posner, Richard A., "A Statistical Study of Antitrust Enforcement," *Journal of Law and Economics* 19 (Oct. 1970), 365–419.

Salop, Steven C., and Lawrence J. White, "Private Antitrust Litigation: An Introduction and Framework," in L. J. White (ed.), *Private Antitrust Litigation* (Cambridge MA: MIT Press, 1988), 3–60.

Stigler, George, "A Theory of Oligopoly," *Journal of Political Economy* 72 (Feb. 1964).

Werden, Gregory J., "Price Fixing and Civil Damages: Setting the Record Straight," Discussion Paper No. 88-13, Economic Analysis Group, Antitrust Division, U.S. Department of Justice, Washington, D.C., dated September 20, 1988.